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Insider power and wage determination in Bulgaria: an econometric investigation

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Keywords *Labour market, Public companies, Private companies, Bulgaria*

Abstract *Using a unique three-digit firm-level data set of all medium- and large-sized manufacturing enterprises in Bulgaria covering the years 1997/1998, and investigation is conducted into how wage determination is related to ownership status. Building on a slightly modified version of the right-to-manage model, the pooled OLS, panel and first-difference TSLS estimates show statistically significant differences in the share of rents taken by workers employed in state, private domestic and foreign firms. Taking account of firm heterogeneity, it is found that rent sharing is nearly non-existent in foreign-owned firms, while the level of pay is higher compared with state-owned companies. Further, rent sharing seems to be highly pronounced in state-owned enterprises, while on average domestically private-owned companies are characterised by less rent sharing. Overall, the robustness checks confirm these findings.*

List of variables

Avg_wage	Running average regional real wage
FOR	Ownership intercept variable: fraction of shares held by foreign owners
FORprof_L	Real profits per employee interacted with the foreign ownership variable
FORsales_L	Real sales per employee interacted with the foreign ownership variable
FORva_L	Real value added per employee interacted with the foreign ownership variable
FORva_L_{t-2}	2-period lagged value of FORva_L
FORva_L_{t-3}	3-period lagged value of FORva_L
FORwage_L	Real wage per employee interacted with the foreign ownership variable

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FORwage_L_{t-2}	2-period lagged value of FORwage_L
FORwage_L_{t-3}	3-period lagged value of FORwage_L
majFOR	Ownership intercept dummy variable: value = 1 if fraction of shares held by foreign owners > 50 per cent
majPRIVD	Ownership intercept dummy variable: value = 1 if fraction of shares held by private domestic owners > 50 per cent
PRIVD	Ownership intercept variable: fraction of shares held by private domestic owners
PRIVDprof_L	Real profits per employee interacted with the private domestic ownership variable
PRIVDsales_L	Real sales per employee interacted with the private domestic ownership variable
PRIVDva_L	Real value added per employee interacted with the private domestic ownership variable
PRIVDva_L_{t-2}	2-period lagged value of PRIVDva_L
PRIVDva_L_{t-3}	3-period lagged value of PRIVDva_L
PRIVDwage_L	Real wage per employee interacted with the private domestic ownership variable
PRIVDwage_L_{t-2}	2-period lagged value of PRIVDwage_L
PRIVDwage_L_{t-3}	3-period lagged value of PRIVDwage_L
Prof_L	Real accounting profits per employee
Reg_u	Regional unemployment rate
Sales_L	Real sales per employee
Sector_wage	Average two-digit sectoral real wage
Va_L	Real value added per employee
Va_L_{t-2}	2-period lagged value of Va_L
Va_L_{t-3}	3-period lagged value of Va_L
Wage_L	Real wage per employee
Wage_L_{t-2}	2-period lagged value of wage_L
Wage_L_{t-3}	3-period lagged value of wage_L
Year 1997	Time dummy: value = 1 if year = 1997

Introduction

Following Slichter's (1950) prominent attack on traditional analysis, labour economists have devoted much effort to test for labour market imperfections in the USA and in Western Europe. In contrast, testing for non-competitive elements in the labour markets of post-communist countries has largely been an unexplored field (excepting Grosfeld and Nivet, 1999 [Poland]; Luke and Schaffer, 1999 [Russia]).

The literature offers three types of rent-sharing models in which the firm's ability to pay explains its wage payments: the modified competitive model, the optimal labour contract model and the rent-sharing bargaining model[1].

Empirically, there are different approaches to estimate the effect of industry or firm performance on wages within a collective bargaining framework. First, some studies concentrate explicitly on the impact of profits per employee on the level of pay (Abowd and Lemieux, 1993; Blanchflower *et al.*, 1996; Goos and Konings, 2001). A second group focuses on the effect of value added or sales per employee together with quasi-rents. Lever and Marquering (1996), Nickell and Kong (1992) and Teulings and Hartog (1998) follow this line of reasoning. A third category estimates directly the effect of product market power on wages. For example, Budd and Slaughter (2000), Konings and Walsh (1994) and Van Reenen (1996) provide evidence that a highly oligopolised market structure may be conducive to rent sharing, hence leading to higher wages.

Considering a range of estimation methods, all these studies strongly support the rent-sharing hypothesis. When using the current or the lagged value of the variables measuring firm/sector performance, the elasticity of wages with respect to the internal variable lies in general between 0.01 and 0.09. Using Lester's (1952) methodology for parameterising the size of rent-sharing in labour markets, it is found that 18-36 per cent of wage inequality may be the result of rent sharing. Instrumenting the profit terms by variables capturing truly exogenous profit shocks increases the importance of the rent-sharing parameter greatly. For example, Abowd and Lemieux (1993), using a foreign competition measure, and Van Reenen (1996), using innovation measures as instruments, found a profitability-wages elasticity of 0.27-0.39[2].

In this paper, we concentrate on imperfect competition in the labour market in post-Communist Europe. To this end, we focus on the economics of wage determination in Bulgaria, a reforming country which can be considered as lagging behind in the transition process.

In particular, the paper aims at analysing whether labour market imperfections differ between state, private domestic and foreign companies. By investigating ownership effects on wages, this research belongs to the empirical literature focusing on the effect of ownership on firm performance. The motivation is that the prevailing wage-setting mechanism in state-owned, private domestic-owned and foreign-owned firms reveals to some extent the objectives underlying firms' behaviour. It may reflect the relative strength of their insiders and the relative importance given to profitability considerations.

The analysis draws upon a unique representative panel of firms in manufacturing with detailed information on output and input factors and firm ownership (company accounts data) for the period 1997-1998.

In the remainder, the paper will first discuss the institutional context of wage bargaining in Bulgaria during the transition period. In the next section we describe the model. Then, we present the data set and some summary statistics. The following section discusses the estimation method and confronts the different theoretical hypotheses with Bulgarian firm-level data. Next, we report some robustness checks. The penultimate section gives an interpretation of the results. The final section concludes.

Institutional background

Under central planning, collective bargaining was absent and wage levels and structures were determined by central planning authorities without union input. Trade unions acted merely as workplace representatives of the Communist Party in the state-owned enterprises. Their primary role was to facilitate the achievement of central planners' production goals. Their preferential treatment of union members in terms of the distribution of social benefits produced very high union membership rates. Overall, the emphasis on production enhancement rather than on workers' representation resulted in considerable job dissatisfaction at the onset of transition (Blanchflower and Freeman, 1997).

As wage competition was constrained by nation-wide administrative wage structures applicable in all state-owned firms, managers competed for workers by offering special benefits. This occurred sometimes through a bargaining process with the state for subsidies or other preferential treatment (Flanagan, 1998).

The political and economic transition since the late 1980s has resulted in rapid institutional reform. All countries have developed labour codes stimulating collective bargaining and permitting strikes. The political independent unions have been formed through reform of pre-existing Communist trade unions and entirely new union organisations. For example, a federation of alternative unions, Prodkrepa Confederation of Labour, was established in Bulgaria in 1989. Being anti-communist, Prodkrepa supported free-market economic policies. In early 1990, the Bulgarian Communist unions were transformed into an independent Confederation of Independent Trade Unions (CITUB). Union membership rates have declined from nearly 100 per cent under Communism to 35-60 per cent nowadays. The growth of the small company sector, the increasing number of foreign-owned firms and the increasing use of fixed term labour contracts have undermined the organisation and the representation of workers by unions. Their representation is primarily confined to current and former state-owned enterprises, a declining sector characterised by less advantageous key working conditions compared with the expanding private sector (Cox and Mason, 2000).

Despite the pluralisation of capitalism, tripartite institutions – referring to social dialogue among governments, reformed and alternative unions, and employers' organisations – have appeared at the beginning of the 1990s in all formerly Communist countries. In line with the requirements of a new labour code, making social dialogue and social partnership mandatory, the National Council for Tripartite Cooperation emerged in Bulgaria at the beginning of 1993. Tripartite negotiations are centred on developments in the area of industrial relations and labour legislation, public-policy formulation and prevention and settlement of social conflicts (Iankova, 1998).

The development of tripartism has led to a two-tier bargaining structure in all countries. Although systems of wage determination differ between countries, they are a combination of national-level decision making and collective bargaining or unilateral management determination[3], the balance varying between countries. In contrast with most of the countries, union power is relatively large and wage determination occurs through bargaining in Bulgaria and Poland (Martin and Cristescu-Martin, 1999). This institutional feature motivates our choice of Bulgaria for analysing labour market imperfections at the firm level.

At the central level, social partners generally determine a national minimum wage and negotiate wage growth norms. General economic pressure and industrial restructuring have called for enterprise-specific solutions. Consequently, by the second half of the 1990s, collective bargaining has become, *de facto*, an enterprise-level activity. Legally, such bargaining is expected to respect legislation and to occur within the tripartite framework. As relative wage changes are needed to provide incentives for skill acquisition and further sectoral redeployment of labour, decentralised bargaining is likely to stimulate better long-run economic performance in these countries.

Theoretical and empirical framework

Theoretical framework

In accordance with the wage determination system applicable to Bulgaria, as described earlier, wages are considered to be the result of bargaining between the unions and the firm represented by its manager.

Right-to-manage model. In attempting to model how labour unions/insiders may raise their wage above the outside option or the alternative wage, the literature developed two basic models of rent sharing in a collective bargaining framework: the right-to-manage model (Nickell and Andrews, 1983) and the efficient bargaining model (McDonald and Solow, 1981)[4]. As both models produce the same wage equations, we rely on the right-to-manage model[5].

In its general form, the right-to-manage model has a generalised Nash bargaining solution which is the product of each agent's gain from reaching an agreement, weighted by their respective bargaining strengths. The firm's net gain is equal to its profit function $\pi = Y(N) - WN$, where $\pi =$ real profits, $W =$ the real wage, $N =$ the employment level and $Y(N) =$ a twice differentiable real value added function characterised by $Y'(N) > 0$, $Y''(N) < 0$ and $Y(0) = 0$. Using the expected utility representation, the union's net gain amounts to $[u(W) - u(A)](N/L)$, where $L =$ membership, $A =$ the outside option, $W > A$ and $0 < N \leq L$.

When β represents the bargaining strength of the employees, the generalised Nash bargaining problem can be written as:

$$\begin{gathered}\max_W \Phi = \{U[W, N(W), A]\}^{\beta}\{\pi[W, N(W)]\}^{1-\beta} \\ \Downarrow \\ \max_W \Phi = \{\tfrac{N}{L}[u(W) - u(A)]\}^{\beta}\{Y(N) - WN\}^{1-\beta}\end{gathered} \tag{1.1}$$

Modified right-to-manage model. What is the impact of labour market conditions on wage determination at the firm level? Can we observe upward responsiveness of wages to labour productivity? Is this effect differentiated according to insider power? To study these questions, we modify the basic right-to-manage model slightly.

First, union members are regarded as being risk-neutral. Second, given our short-run focus, we assume that employment is not an argument in the union's utility function. In other words, the median worker is considered to be certain of his job such that utility is only affected by the wage rate in the short run. Algebraically, the real wage W is assumed to result from the maximisation of the following Nash-bargaining maximand:

$$\Phi = \{W - A\}^{\beta}\{Y - WN\}^{1-\beta} \tag{1.2}$$

with A the outside option expressed in real terms, π real profits and Y real value added.

Rewriting equation (1.2) in logs, we obtain:

$$\ln \Phi = \beta \ln(W - A) + (1 - \beta)\ln \pi \tag{1.3}$$

For a maximum of this function with respect to the wage, it is required that:

$$\begin{gathered}\beta \frac{\partial \ln(W-A)}{\partial \ln W} + (1 - \beta)\frac{\partial \ln \pi}{\partial \ln W} = 0 \\ \Downarrow \\ \beta \frac{W}{W-A} + (1 - \beta)\frac{\partial \pi}{\partial W}\frac{W}{\pi} = 0\end{gathered} \tag{1.4}$$

Since $\partial \pi / \partial W = -N$, equation (1.4) can be written as:

$$W = A + \frac{\beta}{1-\beta}\frac{\pi}{N} \quad (1.5)$$

Third, value added is used to capture the firm's good fortune. Our motivation is twofold. The first reason is that it can be argued that accounting profits may bear little resemblance to economic profits or the size of rents. This is especially the case when economic and accounting rates of depreciation differ significantly. Hence, the value-added variable might be a better proxy for economic rents. Second reason is that although the profits per worker variable has the advantage that it controls for all costs, it has the disadvantage that it is negatively related to wages by construction, hence creating a severe endogeneity bias. Switching to value added per employee eliminates the direct endogeneity problem[6].

As real profits equal real value added minus the wage bill ($\pi = Y - WN$), the term $(\beta)/(1-\beta)W$ should be added to both sides of equation (1.5) to obtain an expression for the equilibrium wage in function of real value added per worker:

$$W = A' + \beta\frac{Y}{N} \quad (1.6)$$

were $A' = (1-\beta)A$.

According to this model, real firm-level wages are affected by both internal conditions or inside forces (represented by value added per employee) and external factors or outside forces (taken up by the outside option or the alternative wage) and the bargaining power of the employees. Most studies measure the latter parameter explicitly by variables such as union density/coverage or strike intensity. In this study, the correlation between real wages and variables measuring firm performance is taken to be indicative of insider power.

As we are particularly interested in the impact of ownership status on wage formation, we add an extra dimension to the original right-to-manage model. More specifically, the ownership form of the firm is expected to be a potential determinant of insider power. Therefore, in the estimations we allow both for ownership-specific (state-owned, private domestic-owned and foreign-owned) coefficients on the proxy for internal factors and ownership intercept coefficients.

Empirical model

The specification that acts as the bedrock for our regressions is given by:

$$\begin{aligned} W_{it} = {} & \alpha + \alpha_{\mathrm{privd}}\mathrm{PRIVD}_{it} + \alpha_{\mathrm{for}}\mathrm{FOR}_{it} + \delta A_{rt} + \gamma\mathrm{valad}_{it} \\ & + \gamma_{\mathrm{privd}}\mathrm{PRIVD}_{it}\mathrm{valad}_{it} + \gamma_{\mathrm{for}}\mathrm{FOR}_{it}\mathrm{valad}_{it} + \varepsilon_{it} \quad (1.7) \end{aligned}$$

where subscript i is used to index observations on individual firms, r denotes region and t represents year.

The dependent variable is the annual real wage per worker in firm i. In our main results, real value added per worker (valad) is used to capture the internal conditions of the firm. To check robustness, we will later proxy inside forces by the impact of real profits or real sales per worker on real firm-level wages.

The variables PRIVD and FOR refer to the fraction of shares held by private domestic owners and foreign owners in the firm at time t. The ownership category that is left out is the state, which refers to the fraction of shares in the firm held by the state, municipalities or treasury.

The conditions on the labour market (external conditions or indicators of the workers' fall-back utility) are represented by different variables (denoted by A in equation (1.7)). First, we take the running average regional wage per worker as a proxy for the outside option or alternative wage. Second, the regional unemployment rate, which takes into account low mobility of labour, is used. Controlling for region-specific effects is particularly important in the context of Bulgaria as there are considerable disparities between the regions in which the firms are located (see later). Finally, the external factors are captured by the average two-digit sectoral wage, reflecting relative conditions of labour demand and supply. After all, wage bargainers are expected to consider industry-wide wage agreements in their own wages. Next to it, the specifications combine the average sectoral wage with the regional unemployment rate.

A white noise error term is represented by ε. All specifications include a dummy year to capture possible unobservable aggregate shocks. Finally, we control unobserved firm heterogeneity by including a firm-level fixed effect, even within the separate ownership groups.

We prefer to estimate equation (1.7) in levels rather than in logs for two reasons. First, because of the presence of loss-making firms in our data, the use of logs would have necessitated discarding observations from poorly performing firms. This would possibly introduce problems of selection bias. Second, the levels-specification is most consistent with theoretical models underlying the bargaining literature.

Hypotheses. When outside forces are proxied by either the average regional wage or the average sectoral wage per worker, the bargaining framework predicts a positive coefficient of δ. In the case where A is measured by the regional unemployment rate, the sign of δ is expected to be negative since regional unemployment signals labour market tightness to the insiders. The higher the regional unemployment rate, the lower the probability of finding an alternative job in the same region. As a result, the insiders will moderate their wage claims.

Based on the competitive-wage arguments, the level of inside wages in private firms is expected to be higher than the wage level in state firms, i.e. α_{privd} and α_{for} are expected to be positive. Next to it, it is often argued that foreign firms have superior technological know-how and expertise compared with domestic firms, which leads them to be more efficient than domestic companies (Djankov and Hoekman, 1998). In the literature on multinational enterprises (MNE)[7], it is also a stylised fact that foreign firms pay on average higher wages than their domestic counterparts, even controlling for a wide range of worker and/or firm characteristics such as worker occupation and capital intensity. The reasoning behind this is that foreign firms try to attract higher quality workers by rewarding them more. Therefore, we expect the coefficient of α_{for} to be larger than the coefficient of α_{privd}.

A priori, we anticipate an upward responsiveness of real firm-level wages to rents per worker. At the same time, important differences of the insider effect across different types of firms (i.e. differences in the absolute size of the coefficients γ, γ_{privd} and γ_{for}) are expected.

Data and summary statistics

We use a panel of 1,514 manufacturing firms from the 28 regions ("oblasti") in Bulgaria. To be included in the data set at least one of the following criteria has to be satisfied: number of employees greater than 100 or total assets and sales exceed US$8 and 16 million, respectively.

All the variables are taken from published annual company accounts which were made consistent across countries by Bureau Van Dyck. The monetary variables are expressed in millions of leva. The data set is called the Amadeus data set.

Although the data cover the period 1994-1998, we will focus the analysis on the period 1997-1998 because it is only for these two years that we have detailed information on the ownership structure of the firms.

This unique data set allows us to make at least two main contributions. First, scarce existing empirical work on transition economies in this field mostly had to rely on small samples of firms collected through surveys. In contrast, our sample contains virtually the entire population of medium and large firms in manufacturing. Comparing the employment and sales coverage of the data with the total employment and sales in manufacturing reported in the statistical yearbooks reveals that the data cover 82 per cent of total sales and 66 per cent of total employment in manufacturing[8]. In the Appendix, Table AI presents a comparison between the sample and the population for all two-digit sectors. From Table AI, it follows that the data are also quite representative at the two-digit NACE sectoral level. Thus, our data set can be considered as fairly representative of the population of firms operating in the manufacturing sector. Furthermore, the data set is collected from company accounts at the three-digit level of sectoral disaggregation. To our knowledge, this kind of detailed firm-level data for a transition country has not been used before for this purpose.

A second strength of the data set is that it offers a detailed information on the ownership structure of firms for two consecutive years. In particular, we know the fraction of shares held by the state and by private investors and can observe its evolution over time. Next to it, we are able to make a distinction between private domestic investors and foreign investors. Based on this detailed description of the shareholding structure and in line with the literature on the effect of ownership in transition countries, we also classify firms as being majority state-owned, majority private domestic-owned or majority foreign-owned. Earlier studies for Central and Eastern Europe in this field had to rely on ownership dummies to investigate the crucial question of how wage formation is related to the form of ownership (Grosfeld and Nivet, 1999; Luke and Schaffer, 1999). The use of fractions of shares also enables us to perform some additional robustness checks. Table I shows how ownership is distributed on average.

Note from Table I that state ownership is relatively important, but decreased very much in 1998, while the average fraction of shares in 1998 held by private domestic owners increased up to 68 per cent. This ownership change is the result of a major privatisation program that finally took place (EBRD, 2000).

The fraction of shares held by foreign owners is only 4 per cent on average, meaning that only a relatively small fraction of firms do have some foreign participation. If we look at shareholding in foreign firms only, however, we can see that the low average share of foreign ownership hides the fact that foreign investors are concentrated in a few firms. For example, in 1998 119 firms had a foreign owner, the average share was 63 per cent and 83 firms were holding more than 50 per cent of the shares. Hence, foreign investors own most of the time a majority share.

Concentration on shareholding in domestically private firms only reveals that private domestic investors hold on average 80 per cent of the total shares in domestically private firms.

Finally, we can observe that the fraction of private domestic and foreign firms in the total number of firms increases over time[9]. Recently, the flows of foreign direct investment started to increase rapidly. By 1998 there was almost

Table I. Distribution of ownership

	Mean (St. dev.) 1997	Mean (St. dev.) 1998
Per cent state ownership (STATE)	0.34 (0.38)	0.27 (0.35)
Fraction of STATE firms in total number of firms	0.70	0.66
Per cent state of all state firms	0.49 (0.36)	0.40 (0.35)
Per cent private domestic ownership (PRIVD)	0.62 (0.39)	0.68 (0.37)
Fraction of PRIVD firms in total number of firms	0.79	0.83
Per cent private domestic of all private domestic firms	0.78 (0.26)	0.82 (0.23)
Per cent foreign ownership (FOR)	0.04 (0.17)	0.05 (0.19)
Fraction of FOR firms in total number of firms	0.06	0.08
Per cent foreign of all foreign firms	0.68 (0.23)	0.63 (0.29)
Number of majority state firms	332	269
Number of majority private domestic firms	897	1,150
Number of majority foreign firms	63	83

Source: Amadeus Database

a tenfold increase in FDI compared with 1991 (EBRD, 2000). The rising number of firms in total reflects a better coverage in the latest year and indicates that our analysis draws upon an unbalanced panel.

The regional variables (at the NUTS3-regional level) are collected from the National Statistical Institute and the United Nations Development Program. Table II presents summary statistics for the variables used in the regression analysis.

All annual wages are expressed as real wages per worker, i.e. nominal wages deflated by a three-digit producer price index, normalised to 1 in 1995. This price index is obtained from the central statistical offices. Alternative wage 1 represents the running average regional wage per worker and alternative wage 2 the average two-digit sectoral wage per worker. Profits, value added and sales per worker are also expressed in real terms. Value added is calculated as sales minus material costs and profits as value added minus the wage bill. Average measures are constructed by dividing the variables by the average number of employees in each firm for each year, respectively. The average employment level is 361 and employment ranges from six to 16,280 employees.

From Table II, it is clear that profits, value added as well as sales vary much more than wages. Restricting the data to the balanced panel indicates that over the sample period the producer price index increases by 24 per cent and the average employment level decreases by 5 per cent. Real wages per worker, real sales per worker and real value added per worker display an increase of 17, 27 and 36 per cent, respectively, whereas real profits per worker decreases by 7 per cent.

The spread of the regional unemployment rate reveals considerable regional variation and disparities. A closer look at the regional data shows that the

	1997			1998			1997-1998		
Variables	Obs.	Mean	St. dev.	Obs.	Mean	St. dev.	Obs.	Mean	St. dev.
Employment	1,306	374.12	759.47	1,381	348.03	693.74	2,687	360.71	726.41
Average wage	1,043	98.62	101.86	1,109	112.22	76.55	2,152	105.63	89.95
Alternative wage1	1,514	97.71	16.71	1,514	111.81	15.17	3,028	104.76	17.44
Alternative wage2	1,509	97.11	25.53	1,509	110.33	32.87	3,018	103.72	30.16
Regional unemployment rate	1,514	15.42	5.64	1,514	16.39	5.25	3,028	15.91	5.47
Sales per employee	1,045	652.04	1,245.87	1,111	1,349.29	18,863.37	2,156	1,011.34	13,570.30
Profits per employee	1,038	178.08	601.59	1,106	303.45	3,277.26	2,144	242.75	2,391.06
Value added per employee	1,038	277.09	663.89	1,108	415.98	3,279.45	2,146	348.80	2,401.73

Source: Amadeus Database; NSI (1998, 1999); UNDP (1999)

Table II.
Summary statistics

region Sofia (situated in the district Sofia-City) is characterised by the lowest average unemployment rate (8.4 per cent) and the highest real wages. In contrast, the highest unemployment rate is to be found in the region Montana (situated in the North-West, district Montana). Real wages are the lowest in the region Kyusendil (situated in the South-West, Sofia district). The National Human Development Reports (UNDP, 1999, 2000) indicate that the income gap between the richer and the poorer regions is almost 50 per cent.

Table AII in the Appendix presents summary statistics by ownership category. The average employment level is the highest in majority foreign firms (652), followed by majority state firms (441) and the lowest in majority private domestic firms (331) (see lower part of Table AII).

Workers in majority foreign firms get the highest wages (mean wage of 153). Wages in majority state and majority private domestic companies are much lower (mean wage of 100 and 106, respectively).

Privatisation is clearly associated with better firm performance. Majority of private domestic firms outperform majority state firms. Furthermore, majority foreign firms outperform majority state firms as well as majority private domestic firms. Using the same data set, recent empirical research has found that privatisation has positive effects on firm performance and that foreign-owned firms do better than domestically private-owned firms (Estrin *et al.*, 2001). Strikingly, 18 per cent of the majority state companies (87 out of 476) are classified as loss-making firms, reporting non-positive profits per employee over the sample period.

Concentration on the evolution of the variables over time and restriction to the balanced panel reveals that, apart from the employment level and real profits per worker, majority state, majority private domestic and majority foreign companies behave in the same way.

Two proxies for internal conditions, real sales and real value added per worker, show a considerable increase over the sample period. Real value added per worker, for example, increases by 32, 21, 32 and 16 per cent in the total sample, in majority state firms, in majority private domestic enterprises, and in majority foreign companies, respectively. Real profits per worker decrease by 6, 4 and 7 per cent in the total sample, in majority state companies and in majority private domestic firms, while majority foreign companies record an increase in real profits per worker of 12 per cent. Real wages per worker display a rise of 17 per cent in the whole sample and in majority private domestic firms, an increase of 10 per cent in majority state companies and a rise of 25 per cent in majority foreign firms.

Estimation method and results

Estimation method

The estimation strategy consists of three parts. First, in line with the existing empirical research, the pooled ordinary least-squares estimator is used as a benchmark for cross-sectional time-series estimates. As this estimation method ignores the panel structure of the data, the results are only presented for illustrative purposes and are not reported (but are available upon request).

The panel structure of the sample offers the opportunity to control for firm-specific heterogeneity, which may capture various unobservables, such as the quality of capital and labour. Therefore, our principal results are obtained using the panel data estimation method.

In the last part, we try to deal with two problems that have not been addressed so far. First, simultaneity may obscure the true relationship between wages and the variables reflecting internal conditions. Second, the level of employment entering both the definition of the wages and the measure of rents per worker raises the standard problem that measurement error may induce spurious correlation between these two key variables.

In general, there are two approaches to circumvent these problems: the recursive equation approach and the instrumental variable approach. To check the robustness of the results, this paper adopts the instrumental variable approach. Using the ordinary two-stage least-squares method, however, brings us back to the problem we faced when employing pooled OLS, i.e. the panel structure of the data is completely ignored. In order to remove pay differentials between the firms which are not related to firm performance, but to the character of the firm, the first-difference instrumental variable method suggested for dynamic fixed-effects models by Anderson and Hsiao (1982) is used. Under the assumption that endogeneity is constant across the years, these results are expected to be in line with those obtained by the fixed-effects estimator.

Estimation results

The pooled OLS as well as the panel estimates consist of two parts. In the first part, interaction-effects are ignored and the modified right-to-manage model augmented with the variables capturing the fraction of shares held by private domestic and foreign investors is estimated[10]. This part hence analyses to what extent the level of pay is related to ownership status. The pooled OLS estimates do not allow for clear-cut conclusions concerning the level of inside wages in private domestic-owned and foreign-owned companies. Once we control for unobserved fixed effects, however, we do find evidence that private firms pay higher wages than state firms. In accordance with the MNE literature, we also find that foreign firms pay the highest wages.

In the second part, we test whether the ownership structure of the firm is a potential determinant of insider power and, hence, whether labour market imperfections are differentiated according to ownership status. In other words, we look not only at the direct effects of ownership changes, but also at the interactions with firm performance. The benchmark ownership type is state-owned firms. The interactive coefficients on value added for private domestic and foreign firms are deviations from this benchmark.

For all the specifications, Wald tests show that the rent-sharing coefficients are jointly significant at the 5 per cent level. They also reject equality of the rent-sharing coefficients at the 5 per cent level. An explanation of the variables used in the regression analysis is given in the Appendix.

Pooled ordinary least squares. The pooled OLS estimates indicate that workers in state-owned firms seem to appropriate a large part of the rents. On average, the employees' capacity to capture productivity gains is significantly lower in both private domestic and foreign firms.

As expected, foreign firms pay significantly higher wages than the state-owned companies. Finally, the estimates show that outside forces play an important role in the wage determination process.

Panel data estimation. The cross-section evidence shows strong effects from two potential sources of rents on to wages, i.e. (some form of) ownership and value added. It can, however, be objected that such evidence is unconvincing as the cross-section estimates are biased if omitted variables, of which the unobserved quality of workers is probably an important one, are correlated with ownership and value added[11]. Since such unobserved fixed effects are likely to be positively correlated with private ownership, we are implicitly controlling for one of the potential sources of endogeneity of ownership by using the fixed-effects estimator (Estrin *et al.*, 2001).

We control for firm heterogeneity for each individual firm, even within the different ownership categories.

From the first part of Table III it follows that for all specifications the Hausman test indicates that we should rely on the fixed-effects model[12].

With respect to rent sharing, we find again that employees in the state-owned firms manage to cream off a significantly larger share of the rents compared with workers in private companies. Foreign-owned firms are in fact characterised by zero rent sharing. On average, the point estimate for state-owned firms is 0.052, which yields an elasticity of 0.11 at sample means. As the variable measuring the ratio of value added to employment in state-owned companies has a mean value of 205.65 and a maximum value of 4,910.18, the state value-added series on the microeconomic level is much more volatile than that of a typical economic variable. As a result, an elasticity of 0.11 has very important consequences. The standard deviation of value added per employee in state firms equals 389.48 and the range amounts to 5,731.34. This large range is probably due to outlying observations.

There is no generally accepted way to parameterise the size of rent sharing in labour markets but it is possible to calculate a version of Lester's (1952) "range" of wages. Given the large variance of average value added, we assume that the width of the distribution can be thought of as two standard deviations. Under this assumption, the regression estimates allow a tentative calculation of

Constant	14.101	91.077***	−1.707	0.054
	(18.220)	(11.614)	(11.006)	(14.422)
Va_L	0.049***	0.050***	0.055***	0.055***
	(0.010)	(0.010)	(0.010)	(0.010)
PRIVD	20.474**	21.159**	18.461**	18.453**
	(9.301)	(9.391)	(8.927)	(8.931)
FOR	67.318***	67.988***	47.567***	47.576***
	(14.963)	(15.110)	(14.505)	(14.513)
PRIVDva_L	−0.017	−0.017	−0.036***	−0.036***
	(0.013)	(0.013)	(0.012)	(0.012)
FORva_L	−0.044**	−0.044**	−0.054***	−0.054***
	(0.017)	(0.017)	(0.016)	(0.016)
Avg_wage	0.650***	–	–	–
	(0.153)			
Reg_u	–	−0.307	–	−0.103
		(0.573)		(0.545)
Sector_wage	–	–	0.836***	0.835***
			(0.085)	(0.085)
Year 1997	−6.374**	−15.746***	−4.394**	−4.511**
	(2.711)	(1.796)	(1.960)	(2.056)
Hausman test	$\chi^2(7) = 47.8$	$\chi^2(7) = 193.8$	$\chi^2(7) = 76.0$	$\chi^2(8) = 84.6$
N	2,040	2,040	2,040	2,040
R^2	0.183	0.167	0.248	0.248
Constant	19.424	104.797***	4.844	8.211
	(18.403)	(11.583)	(10.931)	(14.422)
Prof_L	0.033***	0.033***	0.040***	0.040***
	(0.011)	(0.011)	(0.010)	(0.010)
PRIVD	20.629**	21.487**	16.235*	16.226*
	(9.001)	(9.100)	(8.619)	(8.623)
FOR	68.902***	69.650***	43.727***	43.723***
	(13.985)	(14.141)	(13.604)	(13.611)
PRIVDprof_L	−0.035**	−0.034**	−0.053***	−0.053***
	(0.014)	(0.014)	(0.013)	(0.013)
FORprof_L	−0.075***	−0.075***	−0.077***	−0.077***
	(0.017)	(0.018)	(0.017)	(0.017)
Avg_wage	0.706***	–	–	–
	(0.155)			
Reg_u	–	−0.428	–	−0.197
		(0.582)		(0.551)
Sector_wage	–	–	0.885***	0.884***
			(0.086)	(0.086)
Year 1997	−5.832**	−16.125***	−3.885**	−4.109**
	(2.751)	(1.824)	(1.986)	(2.082)
Hausman test	$\chi^2(7) = 119.8$	$\chi^2(7) = 329.1$	$\chi^2(7) = 68.0$	$\chi^2(8) = 77.6$
N	2,040	2,040	2,040	2,040
R^2	0.160	0.141	0.232	0.232

(continued)

Table III. Wage equation 1997-1998, dependent variable $wage_{it}$ – PANEL (fixed-effects)

Constant	11.289	71.011***	−20.021*	−19.521
	(17.554)	(11.556)	(11.394)	(14.460)
Sales_L	0.043***	0.044***	0.052***	0.052***
	(0.006)	(0.006)	(0.006)	(0.006)
PRIVD	18.371**	17.985*	21.958**	21.954**
	(9.664)	(9.729)	(9.267)	(9.273)
FOR	61.328***	62.892***	49.666***	49.669***
	(16.376)	(16.479)	(15.742)	(15.751)
PRIVDsales_L	0.003	0.003	−0.013*	−0.013*
	(0.007)	(0.007)	(0.007)	(0.007)
FORsales_L	−0.001	−0.001	−0.014	−0.014
	(0.009)	(0.010)	(0.009)	(0.009)
Avg_wage	0.510***	–	–	–
	(0.146)			
Reg_u	–	−0.233	–	−0.029
		(0.542)		(0.517)
Sector_wage	–	–	0.791***	0.791***
			(0.082)	(0.082)
Year 1997	−8.927***	−16.275***	−5.674***	−5.707***
	(2.579)	(1.698)	(1.862)	(1.530)
Hausman test	$\chi^2(7) = 156.9$	$\chi^2(7) = 166.5$	$\chi^2(7) = 97.6$	$\chi^2(8) = 100.0$
N	2,044	2,044	2,044	2,044
R^2	0.269	0.260	0.329	0.329

Notes: *Significant at 10 per cent; **significant at 5 per cent; ***significant at 1 per cent. Hausman test checks for orthogonality of individual effects and other regressors. Standard errors in parentheses. R^2=R-sq. within

Table III.

the spread of pay induced, other factors being held constant, solely by the dispersion of value added across firms. The "range" of wages due to rent sharing in state-owned firms is then approximately 42 per cent of the mean wage[13].

In contrast, the bargaining power of employees in foreign firms seems too weak to let them push up wages in response to productivity gains. The average point estimate for foreign companies is 0.003, which results in a very small elasticity between wages and rents (0.02). Although not always significant, private domestic firms also seem to be characterised by less rent sharing.

Compared with the benchmark case, employees in both private domestic and foreign firms are paid significantly higher wages with wages being the highest in foreign companies.

In general, labour market conditions appear to be important for wage bargaining. A doubling of the region-average wage rate would push up real wages by 64 per cent and a doubling of the sector-average wage rate would increase real wages by 82 per cent. Regional unemployment exerts a downward pressure on wages but this effect is not significant.

In all specifications, the control variable for overall economic changes during the sample period has a statistically significant negative impact on real wages.

This effect is the result of the economic and political crisis in which Bulgaria entered 1997. After a short and fragile burst of growth in 1994 and 1995, the Bulgarian economy relapsed into deep recession in late 1996 with figures showing a real growth rate of − 11 per cent, an annual inflation rate of 311 per cent and currency depreciation hitting 3,000 per cent. The lack of a consistent economic policy and the imitation of market economy measures are blamed for the 1996 crisis (OECD, 1998). In terms of economic performance as well as reform in general, 1997 was the year of the "second start".

Robustness checks

To investigate whether our estimation results are robust to the use of different variables and estimation techniques, two robustness checks are carried out.

The first one is related to the measurement of internal conditions and ownership status. Following the empirical literature, we estimate all specifications using accounting profits per worker or sales per worker instead of value added per worker. Next to it, we use a slightly different sample to investigate whether majority shareholding has an effect on our earlier results.

The second check refers to the estimation method: the robustness of the fixed-effects estimator is checked by applying the first-difference instrumental variable approach.

For all specifications, Wald tests point to significant differences between the rent-sharing coefficients at the 5 per cent level of significance.

Measurement check

Internal conditions. When value added per worker is substituted by either profits per worker or sales per worker, the pooled OLS estimates are well in line with the earlier results, i.e. rent sharing in private firms is considerably lower than in state firms[14].

Before interpreting the panel estimates, it should be mentioned that, for all variants of the earlier specifications, Hausman's test always rejects the random-effects model.

The panel results with profits per worker as internal variables are presented in the middle part of Table III. As earlier, rent sharing appears to be most prevalent in state firms. Strikingly, the responsiveness of real wages with respect to profits per worker is clearly lower than their responsiveness with respect to value added per worker: the average point estimate of 0.037 yields a profits-wages elasticity of 0.04 (compared with 0.11 using value added). We can point to the direct endogeneity bias as a possible explanation for the lower estimated rent-sharing effect.

Remarkably, the coefficient on both private domestic and foreign rents is found to be significantly negative and highest in absolute value for foreign firms. This finding leads to the strong conclusion that an increase in profitability in private companies results in lower wages.

The conclusions concerning the level of inside wages and labour market conditions equal those of the first panel estimates.

The last part of Table III reports the panel estimates when sales per worker is substituted for value added per worker. As before, wages in state-owned companies appear to be quite responsive to changes in sales (average point estimate of 0.048).

Surprisingly, the slope coefficients related to both ownership categories appear to be insignificant in two specifications, implying that we are unable to detect significant differences with respect to rent sharing in state, private domestic or foreign firms. In line with the first panel results, private companies pay significantly higher wages compared with the benchmark case. The highest level of inside wages is still to be found in foreign firms. The estimates regarding outside forces confirm the previous results.

Ownership variable. As an additional measurement control, we check whether the results are invariant to the choice of the sample.

When substituting ownership dummies for the fraction of shares held by the various ownership categories, we can firmly conclude that our main results remain robust to restraining the analysis to majority ownership groups only[15].

Methodological check

To correct for possible simultaneity between the internal variable and the wage and to allow for firm-specific effects, the results of the first-difference instrumental variable procedure are reported in Table IV.

The various specifications take up the first differences of all variables. As suggested by Arellano (1989), the instruments are taken in levels. More specifically, the 2-period and the 3-period lagged value of value added combined with the 2-period and the 3-period lagged value of real average wages at the firm level are used as instruments to solve for the least squares of the sample's first moment conditions. The choice of the lagged value of real average wages as additional instrument can be motivated by efficiency-wage considerations. According to this channel, higher wages stimulate work effort and hence productivity.

From Table IV, it follows that the findings of the fixed-effects model appear to be quite robust. In line with the panel estimates, the results confirm the existence of crucial differences in insider effect across different types of firms. Interestingly, the conclusion to be drawn from comparing the fixed-effects estimates with the first-difference TSLS estimates is that the extent of rent sharing in state-owned companies is underestimated using an OLS technique. The first-difference TSLS approach points to an average estimate of 0.072 for state firms, resulting in an elasticity at sample means of 0.15. According to the

Table IV. Wage equation 1997-1998, dependent variable $wage_{it}$ – endogeneity and fixed-effects check

	First-difference TSLS				Panel-fixed effects			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Constant	7.833**	18.841***	2.925	3.147	14.101	91.077***	− 1.707	0.054
	(3.583)	(2.419)	(2.575)	(2.713)	(18.220)	(11.614)	(11.006)	(14.422)
Va_L	0.069***	0.069***	0.075***	0.075***	0.049***	0.050***	0.055***	0.055***
	(0.016)	(0.016)	(0.015)	(0.015)	(0.010)	(0.010)	(0.010)	(0.010)
PRIVD	33.908**	34.959**	28.886**	28.773**	20.474**	21.159**	18.461**	18.453**
	(14.472)	(14.668)	(13.612)	(13.619)	(9.301)	(9.391)	(8.927)	(8.931)
FOR	121.021***	121.453***	86.731***	86.670***	67.318***	67.988***	47.567***	47.576***
	(28.550)	(28.931)	(27.074)	(27.088)	(14.963)	(15.110)	(14.505)	(14.513)
PRIVDva_L	−0.088**	−0.088**	−0.104***	−0.103***	−0.017	−0.017	−0.036***	−0.036***
	(0.033)	(0.034)	(0.031)	(0.031)	(0.013)	(0.013)	(0.012)	(0.012)
FORva_L	−0.156***	−0.157***	−0.164***	−0.164***	−0.044**	−0.044**	−0.054***	−0.054***
	(0.050)	(0.050)	(0.047)	(0.047)	(0.017)	(0.017)	(0.016)	(0.016)
Avg_wage	0.712***	–	–	–	0.650***	–	–	–
	(0.191)				(0.153)			
Reg_u	–	− 0.546	–	− 0.185	–	− 0.307	–	− 0.103
		(0.763)		(0.711)		(0.573)		(0.545)
Sector_wage	–	–	1.146***	1.144***	–	–	0.836***	0.835***
			(0.111)	(0.111)			(0.085)	(0.085)
Year 1997	–	–	–	–	− 6.374**	− 15.746***	− 4.394**	− 4.511**
					(2.711)	(1.796)	(1.960)	(2.056)
Hausman test	–	–	–	–	$\chi^2(7) = 47.8$	$\chi^2(7) = 193.8$	$\chi^2(7) = 76.0$	$\chi^2(8) = 84.6$
N	613	613	613	613	2,040	2,040	2,040	2,040
R^2	0.053	0.031	0.163	0.163	0.183	0.167	0. 248	0.248

Notes: *Significant at 10 per cent; **significant at 5 per cent ***significant at 1 per cent
TSLS difference equations: Standard errors in parentheses – all variables are in first differences, the instruments are in levels. The instruments for all specifications are: $(Va\_L\ PRIVDva\_L\ FORva\_L)_{t\text{-}2,\ t\text{-}3}$ $(Wage\_L\ PRIVDwage\_L\ FORwage\_L)_{t\text{-}2,\ t\text{-}3}$; FE estimator: Standard errors in parentheses – Hausman test checks for orthogonality of individual effects and other regressors. $R^2 = R$-sq within

panel results, the corresponding average point estimate is 0.052 and the corresponding elasticity amounts to 0.11.

Consistent with the panel estimates, in foreign firms the bargaining strength of insiders seems to be too weak to let them push up wages in response to firm performance. Even stronger, the coefficient on foreign rents is found to be significantly negative. In the first-difference TSLS estimates, we find the same result for private domestic firms.

In accordance with the panel estimates, inside wages in private firms are higher than those in state firms.

Interpretation

Our results reveal that labour market imperfections differ between ownership categories, hence, providing evidence of significant ownership effects on wages.

Rent sharing is highly pronounced in state-owned firms whereas both private domestic and foreign companies are characterised considerably by less rent sharing. The results regarding the state-owned and domestically private-owned firms are in line with those of scarce existing empirical research for Poland (period 1992-1994) and Russia (1996-1997) in this field (Grosfeld and Nivet, 1999; Luke and Schaffer, 1999). Both studies use ownership dummies to discriminate between state, privatised and commercialised enterprises and find that the share of rents taken by workers in privatised companies is significantly less than the share taken by state-owned employees.

When focusing on the preferred estimation method, which takes account of firm-specific effects, and when correcting for endogeneity, we find in addition that the level of inside wages is on average higher in private firms compared with state firms and that foreign enterprises pay the highest wages.

Within the bargaining framework, the findings for foreign-owned companies can be explained in two ways. First, foreign firms appear to be characterised by a relatively high level of inside wages, which prevents insiders from translating productivity gains into wage increases. Moreover, if one thinks about a two-stage game in which the location decision of foreign firms occurs after firms and insiders bargain over wages, the "threat of relocation" possibility of foreign firms *vis-à-vis* the insiders increases the bargaining position of foreign firms. If bargaining breaks down, the conflict pay-off (or outside option for the firm) is positive as foreign firms can relocate activity to other countries. Hence, firm mobility curbs insider power. This leads to a low – in our case sometimes a reverse – responsiveness of real wages to productivity gains, generating in turn lower wages in equilibrium (Zhao, 1995). A second possible explanation is that foreign firms are typified by "low"-wage considerations, meaning that they resist letting wages increase in response to productivity improvements.

Next to it, it can be said that wage determination in foreign firms is driven more by profitability considerations than wage formation in state-owned firms. Given the economic and financial crisis in 1996, it is very reasonable to assume that during the considered time-span firms in Bulgarian regions financed most of their investment from retained earnings. The statistically significant negative coefficient on the interaction term related to foreign ownership in the first-difference TSLS estimates may reflect that investment effort strongly dominates other objectives in foreign companies. Higher investment spending in foreign firms could then be explained by more efficient corporate governance and a probably different time horizon underlying economic decisions. Empirical research has already found evidence of investment behaviour being related to form of ownership (Grosfeld and Nivet, 1997).

Finally, the rejection of rent sharing in a non-standard competitive framework may be the result of a clustering of foreign firms in sectors operating in a perfectly competitive market. To gain a first insight into that issue, we divided the sample into six subsamples according to the two-digit NACE industry classification[16]. The underlying idea was to examine in which subsectors rent sharing is not an issue and then to check the homogeneity/heterogeneity of firms according to ownership type within these subsectors. The presence of predominantly foreign-owned firms in subsectors with perfectly competitive elements would then explain our findings. A closer look at the data of the subsamples leads to the following conclusions. First, the different ownership categories are evenly distributed across subsectors. Within each subsector, on average 20-30 per cent of the firms are majority state-owned, 65-80 per cent are majority private domestic-owned and 3-10 per cent are majority foreign-owned. Second, the shareholding structure is quite similar within each subsector. The average fraction of shares held by state owners is 32 per cent, private domestic investors retain on average 64 per cent of the shares and foreign investors hold on average 4 per cent of the shares. Third, within each subsector foreign investors are concentrated in a few firms and hence own most of the time a majority stake. Based on this information, we do not expect to see crucial differences in labour market performance across subsectors. To test this hypothesis, we performed the regressions for each subsector separately. As anticipated, the pooled OLS and the panel results do not point to significant differences in labour market performance across subsectors.

The result that the strongest positive relationship between the firm's ability-to-pay and wages is found in the state-owned firms can be partly explained by the fact that insiders in these companies still play an important role. This is, however, not a sufficient explanation, as increased product market competition (resulting for example, from higher FDI) may prevent insiders from exploiting their power on the bargaining table.

In this context, further research aims at linking explicitly product market performance to labour market performance, hence introducing imperfections in both the product and the labour market. The motivation is that competitive pressure may result in disciplining firm behaviour and that the effect may depend on the ownership type of the firm. Empirical research has already found evidence of the existence of complementarities between competitive pressure and ownership changes in the Bulgarian product market (Estrin *et al.*, 2001).

The economic chaos following the 1996 crisis and the coefficient of the regional unemployment rate showing the expected sign, but being insignificant (in contrast with the other variables capturing outside forces) bring us to another, rather tentative explanation. More specifically, a second interpretation rests on the wage-unemployment ratchet effect, introduced by Lindbeck and Snower (1987). The underlying idea is that, after an adverse shock to which firms reacted by reducing employment, the probability for each employee of being laid off decreases and the incumbent workforce with important bargaining power responds by pushing for higher wages. Based on our evidence, it seems that in the aftermath of the 1996 crisis this effect was really pronounced in the state-owned firms.

Finally, a caveat to our results is the possibility of residual selection bias. It could be that some categories of owners were able to obtain shares in better firms, in ways which are unobservable to the researcher, but possibly observable to the buyers. This problem arises in all studies of privatisation and firm performance.

In our analysis, we argue that the fixed-effects estimator controls for ownership endogeneity. This is valid if the unobservable quality is fixed for each firm. The effect may be dynamic, however, if, for example, the unobservable quality relates to potential for restructuring and improvements in productivity rather than being intertemporally fixed. We control implicitly for this dynamic effect by using the first-difference TSLS method. Nevertheless, the possibility of selection bias should be borne in mind in interpreting our findings.

Conclusion

This paper focuses on wage determination in Bulgaria. Wage equations derived from a slightly modified right-to-manage model are estimated using a unique firm-level data set covering all medium- and large-sized manufacturing firms from the 28 regions of Bulgaria during the period 1997-1998.

Starting from the simplest method to analyse cross-section time-series data, we apply alternative panel procedures, exploiting the rich panel structure of the data as well as controlling for the endogeneity bias.

The central question investigated is whether and how wage determination is related to ownership status. The broad conclusion to be drawn from this paper is that the wage formation process is clearly related to form of ownership: there are statistically significant differences in the share of surplus taken by employees in state, private domestic and foreign companies.

When considering firm heterogeneity, it is found that rent sharing is nearly non-existent in foreign-owned firms, while the level of pay is higher compared with the benchmark, state-owned firms.

Rent sharing appears to be highly pronounced in state-owned firms and rent sharing is on an average considerably lower in domestically private-owned firms. Private domestic firms also pay higher wages than state firms, but the highest level of inside wages is to be found in foreign firms.

Notes

1. For a theoretical framework, see Blanchflower *et al.* (1996).
2. These findings are, however, criticised by some authors who point to the implausibly high induced Lester's range of wages. Moreover, Blanchflower *et al.* (1996) conclude that using sector-average cost shocks to energy as instruments yields the same results as using the lag procedure, although the former estimates are less precise.
3. In many countries, union influence at the enterprise level is limited. Wages are determined unilaterally by management, taking into account labour and product market factors, the minimum wage and the level of earnings needed to avoid union involvement (Martin and Cristescu-Martin, 2001).
4. For a good summary of the issues, see Booth (1995).
5. Another argument for relying on the right-to-manage model is the observation that firms make continuous adjustments to the size of their labour force without any intensive bargaining with unions/insiders except when the adjustments involve compulsory redundancy. Therefore, the right-to-manage model appears to be a more accurate description of reality than the Efficient Bargaining model (Nickell and Andrews, 1983).
6. This, however, does not imply that endogeneity is not an issue anymore. For example, wage shocks affecting productivity may cause endogeneity problems when using real value added per employee.
7. The literature on multinational enterprises (MNE) mainly investigates the effect of foreign-owned firms on the domestic economic environment.
8. Sales coverage ratio = total sales of firms in Amadeus in 1998 divided by total national sales as reported by the National Statistical Offices. The same for employment.
9. Note that the sum of the fractions of, respectively, state, private domestic and foreign firms in the total number of firms does not add up to 1 as each firm can have multiple owners.
10. These results are not reported but are available upon request.
11. Note that the rent-sharing coefficient is subject to three sources of omitted variable bias, i.e. observable worker characteristics (such as variables related to the human capital stock of the workforce), unobservable characteristics of the workforce (variations in the composition of the workforce) and fixed unobservable firm characteristics (time-invariant unobserved heterogeneity in the production processes or work organisation at the firm level). These omitted variables may induce a correlation between the value added per worker term and the disturbance term of the model that could bias the estimates of the rent-sharing effect. For example, the absence of human capital-related variables may drive the observed relation between value added and wages as these variables affect simultaneously the firm's value added (by improving productivity) and wages (since these characteristics have a value on the labour market). As we lack information on the characteristics of the individuals, we cannot control for individual worker characteristics. Having a panel in the firm dimension on value

added, wages and observables, we can, however, correct for fixed firm-specific unobserved heterogeneity by estimating the model with fixed firm effects.

12. A criticism of the use of within-group estimation is that the assumption of non-zero correlation between the time-invariant fixed effect and the exogenous variables does not allow for doing out-of sample inference (Baltagi, 1995). Since we rely on a large representative sample of manufacturing firms, however, we argue that this criticism does not apply to our results.
13. This number emerges from0.11*2*(sd(SOEva_L)/mean(SOEva_L)) = 0.11*2*(389.48/205.65), with SOEva_L real average value added in state-owned companies. Note that this calculation is a modified version of Lester's range of feasible wages. He assumed that – under a normal distribution – the width of a distribution is equal to four standard deviations.
14. Results not reported, but available upon request.
15. Instead of using continuous shareholding variables ranging between 0 and 1, we now replace them with dummies for majority ownership. These results, which are very well in line with the earlier findings, are available upon request.
16. Sectors with a small number of observations were dropped. To be specific, subsample 1 comprises firms from industry 15, subsample 2 from industries 17 and 18, subsample 3 from industries 24 and 26, subsample 4 from industries 27 and 28, subsample 5 from industries 29 and 31 and subsample 6 from industries 36, respectively. Subsamples 1 and 2 both cover 25 per cent of the sample, subsamples 3 and 4 both cover 12 per cent and subsamples 5 and 6 cover 19 and 7 per cent, respectively. In the aggregate, the six subsamples cover about 80 per cent of the original sample. In line with the original sample, 21 per cent of the firms are majority state-owned, 74 per cent majority domestically private-owned and 5 per cent majority foreign-owned.

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Appendix

Table AI. Sales industry coverage using Amadeus data set, 1998

Code	Name	Coverage (per cent)
15	Food and beverages	51
16	Tobacco	80
17	Textiles	100
18	Wearing apparel, fur	44
19	Leather, luggage and footwear	57
20	Wood, straw and plaiting materials	51
21	Pulp, paper and paper products	76
22	Publishing, printing and media	55
23	Coke, refined petroleum products and nuclear fuel	–
24	Chemicals and chemical products	98
25	Rubber and plastic products	55
26	Other non-metallic mineral products	74
27	Basic metals	100
28	Fabricated metal products	56
29	Machinery and equipment n.e.c.	64
30	Office machinery and computers	23
31	Electrical machinery and apparatus n.e.c.	100
32	Radio, TV and communication equipment	100
33	Medical, precision and optical instruments	50
34	Motor vehicles, trailers and semi-trailers	67
35	Other transport equipment	87
36	Furniture, manufacturing n.e.c.	43
37	Recycling	–

Note: Sales coverage ratio = total industry sales in Amadeus/total national industry sales according to the two-digit NACE industry classification. Data on national industry sales are not available in sectors 23 and 37

Table AII. Summary statistics by ownership category

Variables	Obs.	Total sample mean (St. dev.)	Obs.	Majority state firms mean (St. dev.)	Obs.	Majority private domestic firms mean (St. dev.)	Obs.	Majority foreign firms mean (St. dev.)
1997								
Employment	1,163	400.4 (799.0)	303	528.6 (1,398.4)	802	335.0 (390.4)	58	635.3 (553.3)
Average wage	933	101.7 (105.6)	265	98.2 (84.2)	620	100.3 (114.3)	48	137.9 (89.0)
Sales per employee	936	633.5 (1,160.5)	265	498.5 (740.2)	623	651.1 (1,295.2)	48	1,150.8 (1,066.1)
Profits per employee	931	167.3 (595.5)	265	111.5 (396.3)	618	179.1 (667.2)	48	323.8 (495.2)
Value added per employee	931	269.1 (663.3)	265	209.7 (446.5)	618	279.6 (742.0)	48	461.7 (545.8)
1998								
Employment	1,371	346.8 (695.0)	236	328.5 (609.3)	1,058	327.8 (708.5)	77	664.1 (685.7)
Average wage	1,102	112.2 (76.7)	211	102.0 (78.0)	828	110.8 (71.5)	63	164.7 (112.2)
Sales per employee	1,104	1,351.1 (18,922.8)	211	381.8 (428.1)	830	1,578.1 (21,809.9)	63	1,606.3 (2,338.0)
Profits per employee	1,099	303.4 (3,287.5)	211	98.5 (271.5)	827	337.2 (3,741.0)	61	553.9 (2,153.0)
Value added per employee	1,101	415.9 (3,289.7)	211	200.6 (304.1)	828	447.6 (3,743.2)	62	725.8 (2,145.3)
1997-1998								
Employment	2,534	371.4 (744.8)	539	441.0 (1,126.8)	1,860	330.9 (592.5)	135	651.7 (630.1)
Average wage	2,035	107.4 (91.2)	476	99.9 (81.5)	1,448	106.3 (92.4)	111	153.1 (103.2)
Sales per employee	2,040	1,021.9 (13,944.3)	476	446.8 (623.6)	1,453	1,180.7 (16,507.8)	111	1,409.3 (1,902.1)
Profits per employee	2,030	241.0 (2,452.7)	476	105.7 (346.3)	1,445	269.6 (2,863.8)	109	452.6 (1,641.7)
Value added per employee	2,032	348.7 (2,463.3)	476	205.6 (389.5)	1,446	375.8 (2,874.2)	110	610.5 (1,649.7)

Note: In Table AII, the sample is restricted to firms which are classified according to majority shareholding. In contrast, the sample in Table II also contains firms which have multiple owners. Consequently, the number of observations in Table AII differs from the number in Table II

Source: Amadeus Database

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